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Population Studies
A Journal of Demography

ISSN: 0032-4728 (Print) 1477-4747 (Online) Journal homepage: http://www.tandfonline.com/loi/rpst20

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To cite this article: Jorge Rodríguez-Vignoli & Francisco Rowe (2018): How is internal migration reshaping metropolitan populations in Latin America? A new method and new evidence, Population Studies, DOI: 10.1080/00324728.2017.1416155

To link to this article: https://doi.org/10.1080/00324728.2017.1416155

Published online: 30 Jan 2018.

Population Studies, 2018
https://doi.org/10.1080/00324728.2017.1416155

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How is internal migration reshaping metropolitan populations in Latin America? A new method and new evidence

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Internal migration is a key driver of patterns of human settlement and socio-economic development, but little is known about its compositional impacts. Exploiting the wide availability of census data, we propose a method to quantify the internal migration impacts on local population structures, and estimate these impacts for eight large Latin American cities. We show that internal migration generally had small feminizing, downgrading educational, and demographic window effects: reducing the local sex ratio, lowering the average years of schooling, and raising the share of working-age population due to an increased young adult population. Over time, a rise in the proportion of males and a drop in the share of the young adult population moving into cities reduced the feminizing and demographic window effects. Concurrently, a rise in the average years of schooling associated with people moving into cities attenuated the downgrading impact of internal migration on local education levels.

Supplementary material for this article is available at: http://dx.doi.org/10.1080/00324728.2017.1416155

Keywords: internal migration; Latin America; socio-demographic composition; spatial impact; migration analysis

[Submitted March 2016; Final version accepted September 2017]

Introduction

Many countries have seen migration replacing fertility and mortality as the main agent of population change. Alongside international mobility, internal migration is now the primary demographic process shaping national patterns of human settlement. It underpins differences in population change and structure across subnational areas. Understanding and determining how internal migration changes the population composition of local areas is critical for responding to housing, healthcare, educational, and transportation needs; delivering more accurate population forecasts; and assessing the spatial distribution of skills, knowledge, and labour.

Migration research has focused on understanding the factors that trigger migration. Less progress has been made on quantifying the effects of migration on changing the socio-demographic composition of local areas (Rowe, Bell, et al. 2017). This dearth can be traced partly to the absence of a comprehensive methodological approach to estimate these effects. Prior work has typically used three sets of approaches to quantify the spatial impacts of migration: comparative socio-demographic profiles, net migration-based measures, and population growth equations. However, these approaches have failed to quantify the migration impact of multiple population subgroups into a single indicator effectively, and do not assess the impact of migration on a wide range of socio-economic indices, such as the dependency ratio or Gini coefficient, at a fine geographical scale.

To redress these limitations, the research described in this paper aimed to develop a new method to estimate the impacts of internal migration on the socio-demographic composition of local areas, and to quantify these impacts on eight large Latin American (LA) cities in Ecuador, Panama, and Mexico. These cities were selected because of their importance to their national urban, migration, and economic systems (ECLAC 2012) and because we argue they provide valuable insights into the impacts of internal migration on the population composition of areas within LA countries.

The proposed method relies on census-based migration matrices, which provide information on local populations at the census date and some earlier year. In the context of our method, information at the census date is considered to capture the spatial distribution of population attributes after internal migration has occurred, and information at some earlier year is used to represent a hypothetical scenario of no internal migration. Given the reliance on census data, a key assumption of the method is the time invariability of population attributes. While this assumption is reasonable for time-invariant population attributes (e.g., sex) and those that change in a predictable way over time (e.g., age), it is less adequate for time-varying attributes (e.g., employment status). Our method, however, has the advantage of effectively isolating the impacts of internal migration by removing the effects of international migration, births, and deaths, and providing a way to decompose the overall change in the socio-demographic composition of areas due to internal migration into the effects from in-migration and those from out-migration. It thus provides an effective tool for progressing our understanding of the relative contributions of in- and out-migration to overall change in the socio-demographic composition of areas.

The paper is structured as follows. The next section reviews the literature on patterns of internal migration in LA countries and highlights existing theoretical discussion on the impacts of these movements on the socio-demographic composition of large cities. We then discuss the key limitations of existing approaches for quantifying internal migration impacts. In the 'Methodological framework' section, we propose a method that addresses these shortcomings and provides a summary statistic that exploits the use of census data. The following section discusses our estimates of the impact of internal migration on the age, sex, and educational population compositions of large cities in Panama, Mexico, and Ecuador. Finally, we summarize our main conclusions and suggest ways in which the proposed method could expand internal migration research.

Background

The internal migration systems in LA countries have experienced major transformations over the last century. Between the 1930s and 1970s, rural-to-urban migration dominated national migration patterns, spurring significant population redistribution (Firebaugh 1979). Fostered by the introduction of import substitution policies after the Second World War (Brea 2003; Rowe 2013a), net rural-to-urban migration accelerated the urbanization process in LA countries, accounting for over 45 per cent of urban growth between 1950 and 1970 (Lattes 1995; Lattes et al. 2004). Urban growth was concentrated in a few urban centres, particularly in the largest cities, resulting in a pronounced population imbalance between the primate city and the rest of each country.

Rural-to-urban migration during the 1930s to 1970s was characterized as a two-step process: first, moves from rural areas to small towns, and then moves from small towns to urban areas (Herrick 1965). Relative to the destination population, these flows were driven by large proportions of young people, females, and the less educated (Elizaga 1972; Herrera 2013). Like in European and North American countries, the out-migration of young individuals and less educated people from rural areas was underpinned by a lack of educational and employment opportunities (ECLAC 2012), but out-migration flows were reported to be more geographically concentrated, being directed to a limited number of urban centres (Zlotnik 1994), reflecting the prominent concentration of service provision in national capitals (ECLAC 2012).

The predominance of females in rural-to-urban migration was a result of the greater migration propensities of women than men in LA countries, with a ratio of three women for every two men moving to cities during the 1940s to 1960s (Simmons et al. 1978; Lawson 1998). This gender selectivity reflected the patriarchal society and a growing urban service economy (Germani 1971; Gilbert 1974; Jelin 1977). Reflecting patriarchal values, some women migrated as part of the family unit, following the male head of household. But, unlike in Africa, large numbers moved to cities to take up informal service jobs, as men dominated agricultural employment and job opportunities in rural areas (Elizaga 1966; Hugo 1993). The emergence and concentration of affluent classes in cities increased the demand for domestic service workers, who were primarily rural in-migrant women (Elizaga and Macisco 1975; Szasz 1995; Chant 1999).

Given the large volume of rural-to-urban migration characterizing the 1930s to 1970s period, LA scholars have long argued that the selectivity of migration flows has shaped the socio-demographic composition of large cities in the region (Villa and Rodríguez 1998; Rodríguez and Busso 2009). While its quantification has been prevented by lack of appropriate data, rural-to-urban migration is claimed to have generated three main effects: (1) a *demographic window* effect, indicated by a rise in the share of working-age population (15–59); (2) a *feminizing* effect, as shown by a decrease in the local sex ratio; and (3) a *downgrading educational* effect, that is, a decline in the local levels of education in large cities.

In contrast, during the period from the mid-1980s to the 2010s, LA countries experienced significant economic and political changes that reshaped internal migration patterns and are likely to have altered the selectivity of internal migration flows into major urban centres (Gilbert 1993). Stimulated by a transition to an open market economic system (based on trade liberalization, privatization, and development of natural resource-based export-oriented activities), the attraction of major LA cities to migrants diminished (Gilbert 1993; Brea 2003; Chavez et al. 2016). Increased foreign direct investment in mining and agriculturally rich regions promoted greater geographical dispersal of employment growth, with large cities reporting increasing out-migration to distant areas (Rodríguez 2011a; Rowe 2014). Between the 1990s and 2000s, capital cities in many countries experienced net migration losses for the first time since the 1900s (Rodríguez 2008, 2011b). This was coupled with a shift in the primary source of in-migrants, with rural areas no longer the main suppliers of migrants to large cities (Rodríguez and Busso 2009). Small and intermediate urban areas are now the primary sources (Rodríguez 2011b), which is an inherent consequence of the high degree of urbanization in LA (United Nations 2015; Bernard et al. 2017).

These shifts in the migration network of LA countries do not appear to have affected the preference of young migrants for large cities (Rodríguez and Busso 2009; Rowe 2013b), but evidence suggests that the sex and educational composition of migration flows into large cities changed (Rodríguez 2004). Internal migration in LA during the 1990s and 2000s appears to have been selective of males and highly educated individuals. Rodríguez (2004) estimated a male-dominated migrant sex ratio and a larger percentage of migrants with university degrees relative to non-migrants in the region of origin, pointing to an over-representation of males and university-educated people in the migration system.

Taken together, these changes suggest that countries' transitions from dominant rural economies to more industrialized, service-based systems appear to have led to a migration network dominated by urban-to-urban migration, with more educated individuals moving to jobs. Internal migration thus appears to have reshaped the socio-demographic composition of LA cities. However, these compositional impacts have not been examined and are yet to be measured and understood. Based on the work reviewed in this section, we conjecture that the over-representation of males and university-educated individuals in the composition of migration flows would have reduced the *feminizing* and *downgrading educational* effects that characterized the 1930s to 1970s period. At the same time, we believe that the continuation of the migration selectivity of young adults continues to have a *demographic window* effect on the population of large cities, increasing the local share of working-age population. The next section reviews the most commonly used existing approaches for measuring the impacts of internal migration, discussing their strengths and limitations.

Approaches to measuring internal migration impacts

A primary impact of migration is redistributing population, contributing to population growth in some areas while leading to declines in others. A crude approximation for measuring the impact of internal migration on local areas involves estimating the magnitude of inflows to a destination and movements in the reverse direction, and quantifying their resulting net balance. Yet, the magnitude of inflows and outflows is not the only dimension that should be considered. To effectively quantify the impacts of internal migration, accounting for origin and destination differentials in migration selectivity is also important: there are significant variations in out-migration rates across population subgroups (i.e., origin differentials). Out-migrants tend to be younger and more educated than the origin population, and more likely to be single and living in rental housing. Similarly, there are systematic variations in in-migration rates across population subgroups, when comparing the in-migrant and non-migrant populations (i.e., destination differentials).

Accounting for these differentials is important in measuring the impacts of internal migration, as they can produce significant compositional changes in origin and destination regions. Migration may influence the local human capital base, accelerate population ageing, and alter the local sex balance. To quantify such compositional impacts, it is important to capture four key population components: (1) the magnitude of in- and out-migration flows; (2) the size of the non-migrant population; (3) the selectivity of migration flows; and (4) the composition of the non-migrant population.

As mentioned earlier, three sets of approaches have been used to quantify the impact of internal migration on the composition of local areas: (1) comparative analysis of socio-demographic profiles; (2) 'population growth equation' approaches; and (3) net migration-based indicators. While useful, these approaches suffer from a series of shortcomings: they fail to provide a single statistical indicator that integrates the four key components, to capture the combined migration impact of multiple population subgroups, or to assess the migration impact on a wide range of socio-economic indicators.

The comparative socio-demographic profiles approach involves analysis of the frequency distributions of in-migrants and non-migrants with respect to a specific characteristic (e.g., Massey and Parr 2012). These distributions are examined to determine the selectivity of migration to a destination region. They provide a visual inspection of migration data, but do not produce a statistical measure that quantifies the impact of internal migration. Moreover, this approach is based on ratios of population to migration inflows, but it does not directly quantify changes in the non-migrant population due to migration outflows. As a result, this approach only captures a part of the changes in local populations caused by internal migration.

Population growth equations provide a more comprehensive approach to quantifying internal migration impacts. This method involves measuring the components of population change: fertility, mortality, and net migration using a cohort survival model (e.g., Green 1994; Gavalas and Simpson 2007). Survival probabilities are applied to derive net internal migration estimates as the residual between population estimates and projections. These net migration estimates are used to assess the impacts of internal migration. This approach, however, does not consider the size or selectivity of non-migrant populations.

Using five-year census transition data, Table 1 illustrates this deficiency. Based on the population aged 5–14, it displays net migration estimates for the city of Quito and the rest of Ecuador, revealing a net migration gain of over 3,600 people in Quito, and a corresponding loss in the rest of Ecuador during the 1996–2001 census period. Based on these balances, misleading conclusions could be drawn—indicating that the scale of the impact of internal migration on Quito and on the rest of Ecuador were of a similar order—if the sizes of the non-migrant populations in Quito and the rest of Ecuador were not considered. However, an examination of net migration rates, which take into account this population component, reveals that the impacts of internal migration for Quito are much larger than for the rest of Ecuador. Thus, while the 'population growth equation' approach provides an idea of the direction of impact of internal migration (i.e., population gain or loss), it does not produce a direct estimate of the resulting change in population composition.

A second limitation of the 'population growth equation' approach is the stringent data requirement of the cohort method. It requires data on births, deaths, internal, and international migration disaggregated by geographic areas. Such data are rarely available in less developed countries. An additional limitation is that this method only returns net migration estimates. This precludes any decomposition of the overall change in population composition in an area into the contributions of in- and out-migration. This is a major constraint to understanding the underlying ways in which internal migration shapes the local population structure.

The third approach is the estimation and comparison of net migration rates based on transition matrices. Net migration rates effectively summarize the overall impact of internal migration flows by

Table 1 Population change and net internal migration for Quito and the rest of Ecuador, over the 1996–2001 census interval, people aged 5–14

Place of residence	Net migration	Population at risk	Non-migrant population	Net migration rate (%)
Quito	3,613	308,389	292,499	1.2
Rest of Ecuador	–3,613	2,384,420	2,364,917	–0.2

Notes: The population at risk is the total population at the start of the census interval (1996). The staying population is the population at risk minus out-migration. The net migration rate is net migration divided by the population at risk.
Source: Authors' elaboration based on five-year migration data from the 2001 Ecuadorian Census.

balancing net region-specific gains and losses due to internal migration (Thomas 1941). Computed for a population subgroup, net migration rates can also offer an assessment of the selectivity effects of internal migration, in addition to its effects on the size and composition of the non-migrant population (e.g., Voss et al. 2001; Champion and Fisher 2003). A negative value for these rates indicates population losses in a particular subgroup due to net out-migration, whereas a positive score points to population gains due to net in-migration.

Net migration rates, however, do not provide a direct estimate of the impact of internal migration on the socio-economic composition of local areas. They only indicate the change experienced by a particular population subgroup as a result of internal migration. Yet, a population subgroup can increase its share of the population at the destination, despite recording net migration losses, if its corresponding net migration rate is lower than that of other population subgroups. For this reason, if we seek to estimate the impact of internal migration on the population composition of an area, we need to compare the net migration rates for all subgroups into which a population can be divided in relation to a particular attribute (e.g., age). For instance, if we seek to assess the impact of internal migration on the sex composition of an area, a comparison of net migration rates for males and females is required. While this approach is useful, its implementation is computationally intensive and complex, as the number of comparisons increases by a combinatorial factor with the number of population subgroups under analysis. For example, if we seek to compare the changes in age composition across eleven age groups, the number of comparisons required would be 55.

The net migration approach suffers from two key additional limitations. First, it does not return a single summary indicator, so the overall impact of internal migration of multiple population subgroups on area cannot be effectively quantified or easily interpreted. Second, the approach cannot be used to estimate the effects of internal migration on population-based socio-demographic indicators, such as average years of schooling or the dependency ratio; measures commonly used by local government agencies and transnational organizations, including the United Nations (UN), to assess the developmental status of areas.

In this paper, we propose a method that produces a summary statistical indicator to quantify the impact of internal migration on the population composition of local areas. The proposed approach overcomes the limitations of existing methods by integrating the four key population components outlined earlier, and exploits the availability of census data, the most commonly available source of internal migration data around the world (Bell et al. 2015).

Methodological framework

Methods

Our proposed approach provides a summary statistical index, the Compositional Impact of Migration (CIM), that quantifies the impact of migration on the socio-demographic composition of local areas. It has been devised to take advantage of census-based origin–destination matrices of migration flows and to capture the interrelated effects of the four key components identified in the previous section. Thus, this approach overcomes some of the key limitations of existing methods for measuring the spatial impact of migration.

The CIM is a counterfactual approach that involves a comparison between a Factual Value (FV) and a Counterfactual Value (CFV). These values are derived from the row and column marginals of a migration matrix based on a statistical indicator and labelled the Migration Impact Indicator (MII) matrix. It is not a matrix of migration flows. Each element in this matrix represents a statistical indicator that measures the socio-demographic composition of the local population, such as the sex ratio or mean years of education. Like in any standard migration matrix, the diagonal elements of the MII matrix relate to the non-migrant population in an area i; off-diagonal elements relate to the migration flow from a region i to a region j; and the row and column marginals relate to the total population in region i at time t, and in region i at the earlier time t–x, respectively.

FVs correspond to the row marginals of the MII matrix, which are based on the population distribution at the census date. Thus, they provide a representation of the socio-demographic structure of regions observed at the census, that is, *after* migration. CFVs correspond to the elements in the column marginals of the MII matrix, which are based on the population distribution at an earlier year: one, five, or ten years previously, as recorded by censuses. Hypothetically, CFVs could be thought of as the expected population composition if there were no internal migration; in other words, if migration had not happened, what would the local socio-demographic structure have been? Subtracting

CFVs from FVs provides a measure of change between the start and end of a census interval and represents our proposed summary statistic, the CIM. The CIM measures the estimated percentage change in the local population structure, as captured by a MII, resulting from net migration redistribution. A positive CIM indicates that internal migration contributed to increase a given MII, for example, the local sex ratio. A negative value denotes that internal migration reduced the MII. The index is computed as:

$$\mathrm{CIM}_i = \mathrm{FV}_i - \mathrm{CFV}_i = \mathrm{MII}_i^t - \mathrm{MII}_i^{t-5}. \quad (1)$$

Using five-year transition census-based migration data, t and t–5 correspond to the census date and five years earlier, respectively. A detailed example of the calculations is provided in the supplementary material.

To create the MII matrix, any statistical indicator can be adopted, including percentages, ratios, averages, or medians, as well as more complex composite metrics, such as the Duncan index of dissimilarity or Gini coefficient. In this section, we use the sex ratio to generate our MII, in order to provide a complete exposition and mathematical formalization of the method. For our analysis in the next section, we also use the share of population by age band and the average years of schooling, to examine the impacts of internal migration on the sex, age, and educational composition.

Using the sex ratio $P(m)/P(f)$ to quantify the impact of internal migration on the local sex composition, the CIM_i is:

$$\mathrm{MII}_i^t - \mathrm{MII}_i^{t-5} = \frac{P(m)_i^t}{P(f)_i^t} - \frac{P(m)_i^{t-5}}{P(f)_i^{t-5}} \quad (2)$$

where $P(m)_i^t$ and $P(f)_i^t$ denote the local male and female populations at the census date (t) in region i; and $P(m)_i^{t-5}$ and $P(f)_i^{t-5}$ represent those populations five years earlier (t–5). Equation (2) can be decomposed into four elements, as shown by Equation (3), to demonstrate that our method effectively accounts for changes in the effect of the four key components that determine the impact of internal migration on the local population composition. These components are: (1) the magnitude of in- and out-migration flows (M_{ij} and M_{ji}); (2) the size of the non-migrant population (P_{ii}); (3) the selectivity of migration flows (M conditional on gender (f and m)); and (4) the composition of the non-migrant population (P conditional on gender):

$$\mathrm{MII}_i^t - \mathrm{MII}_i^{t-5} = \frac{P(m)_{ii} + \sum_{i=1}^{n} M(m)_{ji}}{P(f)_{ii} + \sum_{i=1}^{n} M(f)_{ji}} - \frac{P(m)_{ii} + \sum_{j=1}^{n} M(m)_{ji}}{P(f)_{ii} + \sum_{j=1}^{n} M(f)_{ji}}\text{; where } i \neq j. \quad (3)$$

We can use equation (3) to decompose the overall impact of migration into the impacts of in- and out-migration. The CIM index can be divided into two component indices: an index for inflows ($\mathrm{CIM^I}$) and an index for outflows ($\mathrm{CIM^O}$). The $\mathrm{CIM^I}$ is computed by comparing the MII for a region after migration with the MII for the wider system of migration flows. It captures the migration inflows from every other area in the system and that of the non-migrant population for that region. The $\mathrm{CIM^O}$ is measured by subtracting the MII for a region at the earlier year before the census date—accounting for all outflows to all other zones in the system—from that of the non-migrant population. For the sex ratio, these indices are:

$$\mathrm{CIM}_i^{\mathrm{I}} = \frac{P(m)_{ii} + \sum_{i=1}^{n} M(m)_{ji}}{P(f)_{ii} - \sum_{i=1}^{n} M(f)_{ji}} - \frac{P(m)_{ii}}{P(f)_{ii}} \quad (4)$$

$$\mathrm{CIM}_i^{\mathrm{O}} = \frac{P(m)_{ii}}{P(f)_{ii}} - \frac{P(m)_{ii} + \sum_{j=1}^{n} M(m)_{ji}}{P(f)_{ii} - \sum_{j=1}^{n} M(f)_{ji}}. \quad (5)$$

A key consideration in the implementation of the method is the way in which internal migration is measured. We require a measure that isolates the impact of internal migration. A common problem in defining internal migration relates to the population at risk. The population at the start of a census interval includes both people who die and people who emigrate during the interval. Rates based on this population are confounded by the risks of mortality, emigration, and internal migration. Rees et al. (2000) recommended the use of the population at the end of the census interval, which includes people who were in the country at the start of the census interval, survived, and were enumerated at the census. The resulting count of internal migrants excludes the influences of mortality, emigration, immigration, and fertility. We adopt this definition, with the advantage of effectively isolating the impact of internal migration from other key components of population change. Figure 1 illustrates the age–time classification used in the method; that is, the age at the end of the census interval. For example, it shows that persons who were aged 15–19 at the time of the census and aged 10–14 at the start of the interval are included in

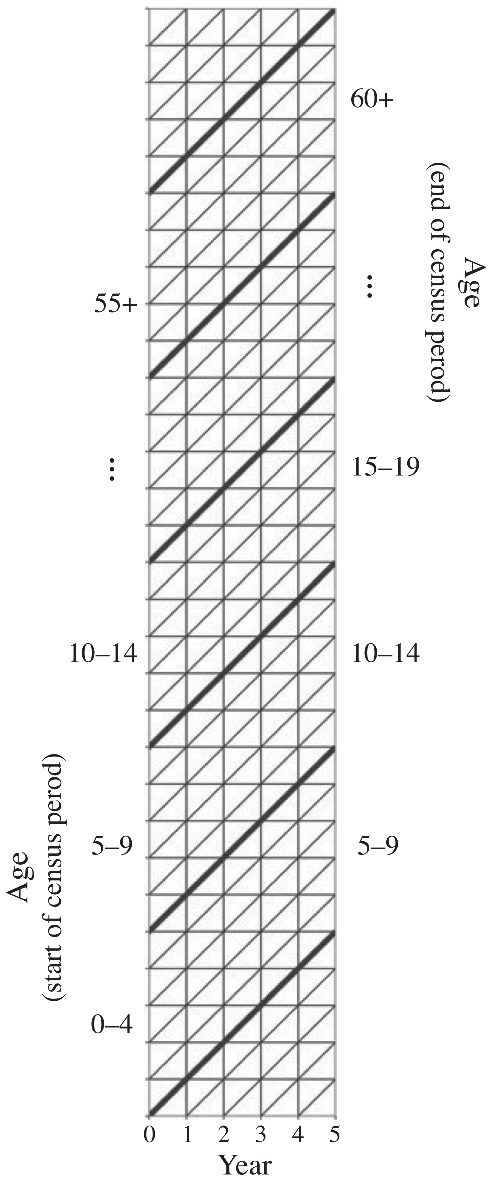

Figure 1 Lexis diagram illustrating age–time observation plan for five-year transition migration data

the calculation, whereas persons aged 0–4 at the census are excluded, as they were not alive at the start of the interval. Figure 1 also shows that persons aged 60+ comprised the group of people who were aged 55+ at the start of the interval. It is important to note that this definition omits moves occurring between census intervals (e.g., return and onward migration).

Another important consideration relates to the way the results from our method must be interpreted. There is a temptation to take a cohort approach, assuming for example that internal migration impact estimates for people aged 5–9 at the end of the census interval are those of the cohort aged 0–4 at the start of the interval. This interpretation is inappropriate for two reasons: first, it adopts a longitudinal view of the impacts of internal migration; and second, it provides the misleading idea that cohort effects are being captured. Our method does not capture cohort effects. It is a counterfactual approach that measures the impacts of internal migration on local areas at the end of a census interval. It compares the factual population distribution in an area at the census date with a counterfactual distribution, representing what the population distribution would have been if a particular population subgroup had not migrated. The results must be interpreted in a 'what if' fashion: for example, what if internal migration of the 5–9 age group had not occurred? What would the age composition of the destination population be? We recognize that people aged 5–9 at the end of the interval correspond to people aged 0–4 at the start of the interval; however, our method measures internal migration impacts at the end of the interval when this population was aged 5–9 at the destination region.

In relation to existing approaches, the proposed method offers four key advantages: (1) it quantifies the internal migration impact on measures of population composition (e.g., gender) in a single index for each; (2) it can be used on a range of socio-demographic statistical measures; (3) it provides an opportunity to contribute to theory development and guide policy design; and (4) it enables us to expand our understanding of structural relationships in the national migration system at fine geographical scales. We elaborate these points in Table 2.

A key limitation of the proposed method is imposed by the assumption of time-invariance of population attributes. Our method relies on a counterfactual distribution, assuming that the observed population characteristics at the census remain stable over time. While this assumption is reasonable for socio-economic characteristics that do not change, or change in a predictable way over time, it is less appropriate for time-varying attributes, such as employment. We note, however, that this limitation is shared by the wide range of analytical approaches based on census data. Despite this, the method has been extensively embraced by migration scholars, as census data only provide information on individual characteristics at the census date.

Data

We applied the proposed methods to measure the impacts of internal migration on the population composition of eight large LA cities: Quito, Guayaquil, and Cuenca in Ecuador; Mexico City, Guadalajara,

Table 2 Key advantages of the proposed CIM method

(1) Captures the internal migration impact of multiple population subgroups in a single index. The proposed method produces a single indicator of the estimated percentage change in the population structure of local areas, resulting from the net migration redistribution of multiple population subgroups. By using the median age or the dependency ratio to build the Migration Impact Indicator (MII) matrix, for example, our method can summarize the overall impact of migration on local age structures. No similar outcome can be achieved through existing approaches. A comparison between net migration rates for multiple population subgroups, for instance, provides a measure of estimated net migration balance for each subgroup, but does not deliver an estimate of how these balances together alter the population structure of local areas.
(2) Allows estimation of internal migration impacts using a wide range of socio-demographic measures. While existing approaches produce migration rates, flows, and percentages to measure the impacts of internal migration (see 'Approaches to measuring internal migration impacts' section), they only provide an indirect estimate of the change in population resulting from internal migration. By contrast, the proposed method produces a direct estimate of the expected percentage change in multiple socio-demographic indicators due to internal migration, such as the sex ratio, the dependency ratio, the average years of schooling, and the share of the population aged 60+. This enables us to make a straightforward assessment of the impacts of internal migration across different dimensions of the local population.
(3) Aids theory and guides policy development. A key additional advantage of the proposed method is its potential to contribute to migration theory and guide policy development. Through the decomposition of the total change in population due to internal migration into the effects of in- and out-migration, the method can be used to determine the leading agent of population change associated with internal population movements. Additionally, through the measurement of the socio-demographic composition of migration, the method can be used to estimate and understand the impacts of migration selectivity. Together, these outcomes represent valuable information for formulating a comprehensive policy framework that targets both in- and out-migration. This is in contrast with existing national policy practices, which focus on influencing in-migration levels (United Nations 2010).
(4) Enables understanding of structural relationships in the national migration system at fine levels of geography. The proposed method uses an indicator matrix (MII) which uses the full origin–destination matrix and allows us to compare the selectivity of internal migration inflows, their counterflows, and the staying population. This comparison produces a representation of the *qualitative* relationships between each of these components by identifying, for example, which of these components is associated with higher sex ratios or higher average years of schooling. Thus, in addition to estimating the overall impact of internal migration on the local population composition, the MII can be used to quantify the impact of migration flows between selected origins and destinations. We do not exploit this advantage in the current paper, as we focus on a small origin–destination migration two-by-two matrix, capturing flows between our chosen cities and the rest of the country.

Monterrey, and Tijuana in Mexico; and Panama City in Panama. The metropolitan areas of each of these cities is home to over 1 million people. We used data from the 2000 and 2010 Census rounds extracted from the online census microdata platform 'REtrieval of DATa for small Areas by Microcomputer' (REDATAM) and city administrative boundaries from the 'Spatial Distribution of Population and Urbanization in Latin America and the Caribbean' database, both hosted by the Latin American and Caribbean Demographic Centre (CELADE). The geographical boundaries used are temporally consistent and correspond to those for the 2010 Census round. These boundaries reflect administrative areas of suburban expansion, accounting for the effects of urban population growth.

We used data on the full population for Ecuador and Panama. For Mexico we drew on data from the extended census questionnaire, equating to a 10 per cent population sample: 10,099,182 and 11,938,402 individuals from the 2000 and 2010 Censuses, respectively. The sampling design for the Mexican extended questionnaire is rigorously tested and samples are carefully assessed following data collection by Mexico's national statistical office (INEGI 2012). Sample weights are available and were applied to make the samples statistically representative of the full census population. We also tested differences in age, educational level, and gender population subgroups between census periods for Mexico. All differences were statistically significant at a 99 per cent confidence level, except for the 45–59 and 60+ age groups in Guadalajara for the analysis of age composition. We believe our results are robust and provide an adequate representation of the way internal migration has contributed to shape large Ecuadorian, Panamanian, and Mexican cities.

The five-year transition data we used to measure internal migration cover the second halves of the 1990s and 2000s (i.e., 1995–2000 and 2005–10). We measured the impact of internal migration on three key population dimensions: sex, age, and educational composition, using three indicators: the sex ratio to measure changes in sex composition; the share of population in five age bands to estimate changes in

Table 3 Statistics used to measure the impact of internal migration on sex, age, and educational composition

Statistic	Formula	Description
Sex ratio	$\frac{P(m)}{P(f)}$	The sex ratio of men per 100 women
Share of population by age band	$\frac{P(a)}{P}$	The share of people in a particular age band (a); where a may be 5–14, 15–29, 30–44, 45–59, or 60+
Average years of schooling	$\frac{\sum s(a)}{P}$	Sum of schooling years of the local population (s) divided by the total population (P); where a indicates an age band and may be 25+, 34–44, or 45–49

age structure; and the average years of schooling for three age groups to quantify changes in educational levels. These indicators comprise our MII (Equation (1)) and are described in Table 3.

As already noted, given our reliance on census data, a key assumption of our method is the time-invariance of population attributes. Information on each individual's situation is only available at the census date and this may differ from their circumstances five years earlier. This creates major difficulties with characteristics such as education that may change over time, especially at young ages, because it is unclear if a rise in the local average years of schooling is the result of (a) the in-migration of highly skilled people; (b) less educated individuals acquiring formal education in the destination after migration; or (c) educational changes in the non-migrant population.

We considered this issue as an integral part of the interpretation and tested the robustness of our results. We measured the internal migration impacts on education for three age groups (Table 3). Consistent with the UN's human development index and educational attainment statistics, we first focused on the average years of schooling for people aged 25+, because it provides a comparable measure across populations and countries (UNDP 2015) and produces more accurate estimates of the impact of internal migration by removing the effect of individuals obtaining education after arrival in the destination. Second, because the in-migration of young and less educated people may still be argued to reduce local average years of schooling at the destination, we conducted a robustness check by comparing the consistency of our results for the 25+ age group with those for two older age groups: 34–44 and 45–49.

Results and discussion

We computed sex, age, and education CIMs by implementing the method outlined in the previous subsection and the measures reported in Table 3. These indices quantify the impact of internal migration and provide empirical evidence on the *feminizing*, *demographic window*, and *downgrading educational* effects that relate to the tentative findings of previous studies on the impacts of internal migration on the population structure of large LA cities. Negative sex and education CIMs would suggest that internal migration had both a *feminizing* and a *downgrading educational* effect on local population structures by reducing the relative share of male relative to female population and the average years of schooling. Coupled with positive age CIMs for the 15–59 population, negative age CIMs for the 5–14 and 60+ populations would suggest that internal migration had a *demographic window* effect by reducing the local dependency ratio, that is, decreasing the share of local population aged 5–14 and 60+ and simultaneously increasing the share of younger people in productive ages 15–59. We also report three additional CIM statistics: the relative CIM, which measures the relative percentage change (rather than the absolute change) in the CIM over the five-year census interval; and the CIM^{I} and CIM^{O}, which quantify the separate contributions of in- and out-migration to the overall impact of migration. The FVs and CFVs used for our calculations are reported in the Appendix (Table A1).

Sex ratios

Table 4 reports the CIMs for the sex ratio. The results show that, except for Cuenca, all cities in the sample display a negative CIM for the 1995–2000 period, suggesting that internal migration operated to reduce the local sex ratio by increasing the share of the female population. These reductions were particularly pronounced in the Ecuadorian cities of Quito and Guayaquil, showing a decrease of 0.7 in the sex ratio. Consistent with previous work (e.g., Elizaga 1966; Alberts 1977), these results indicate

Table 4 Compositional Impact of Migration (CIM): sex ratio, 1995–2000 and 2005–10

	City	Country	Impact indicators			
			CIM	Relative CIM (%)	CIM^{I}	CIM^{O}
1995–2000	Panama City	Panama	−0.29	−0.30	−0.21	−0.08
	Mexico City	Mexico	−0.53	−0.58	−0.42	−0.12
	Monterrey	Mexico	−0.26	−0.26	−0.12	−0.14
	Guadalajara	Mexico	−0.11	−0.12	−0.07	−0.04
	Tijuana	Mexico	−0.30	−0.30	0.07	−0.37
	Quito	Ecuador	−0.70	−0.75	0.07	−0.77
	Guayaquil	Ecuador	−0.71	−0.73	−0.35	−0.36
	Cuenca	Ecuador	1.00	1.16	1.41	−0.41
2005–10	Panama City	Panama	0.02	0.02	0.22	−0.20
	Mexico City	Mexico	−0.24	−0.26	−0.02	−0.22
	Monterrey	Mexico	−0.58	−0.59	−0.20	−0.38
	Guadalajara	Mexico	0.22	0.24	0.40	−0.18
	Tijuana	Mexico	−0.52	−0.53	0.12	−0.64
	Quito	Ecuador	−0.66	−0.71	0.26	−0.92
	Guayaquil	Ecuador	−0.23	−0.23	0.12	−0.34
	Cuenca	Ecuador	0.36	0.40	1.35	−0.99

Note: The CIM statistic indicates the absolute difference between the sex ratio measured at the census date and five years earlier; relative CIM shows the percentage change over the five-year period; and CIM^{I} and CIM^{O} quantify the impacts of in- and out-migration, respectively, on the sex ratio.

Source: Authors' calculations based on census data. Data covering the 1995–2000 and 2005–10 census periods were used for Panama and Mexico, and data covering the 1996–2001 and 2005–10 census periods were used for Ecuador.

that internal migration continued to have a *feminizing* effect on the demographic structure of large LA cities during the second half of the 1990s.

Table 4 also reveals pronounced cross-city differences in the main factor contributing to this *feminizing* effect. While in-migration appears to have been the main contributing force in Panama, Mexico City, and Guadalajara, out-migration was the main driving force in Tijuana and Quito, and both in- and out-migration in Guayaquil and Monterrey. These differences in contribution were paralleled by differences in the direction of their influence, with both acting to shape the reductions in the sex ratio in Panama City, Mexico City, Monterrey, Guadalajara, and Guayaquil, while they operated in opposite directions in Tijuana and Quito. In the latter cities, out-migration appears to have reduced the local sex ratio, while a larger share of males in the in-migration flows—relative to the local non-migrant population—acted to increase the local sex ratio. In absence of this in-migration effect, out-migration would have led to a 0.37 reduction in the sex ratio in Tijuana and a 0.77 reduction in Quito. Thus, a consistent over-representation of men in out-migration flows relative to the non-migrant population has been the main driver underpinning this *feminizing* effect.

In contrast to 1995–2000, the 2005–10 period showed a reduction in or reversal of the feminizing effect of internal migration. Quito, Guayaquil, and Mexico City registered reductions in their corresponding CIMs, from −0.70, −0.71, and −0.53 in 1995–2000 to −0.66, −0.23, and −0.24 in 2005–10, respectively. In contrast, Panama City and Guadalajara saw shifts from a negative to a positive CIM (−0.29 to 0.02 and −0.11 to 0.22, respectively). In Panama City and Guadalajara, these changes point to a major shift in the impact of internal migration on the sex composition, shifting to have a masculinizing effect on the local populations.

As revealed by Table 4, these patterns appear to be largely driven by over-representation of males in in-migration flows relative to the local non-migrant population, offsetting the impact of out-migration, which acted to increase the share of females in the local population. These effects were particularly pronounced in Quito where out-migration reduced the local sex ratio by 0.92. Outflows exceeded an offsetting increase of 0.26 due to in-migration and reduced the local representation of males in the population, leading to an overall reduction in the local sex ratio (−0.66).

Monterrey, Tijuana, and Cuenca represent interesting cases. While the feminizing impact of internal migration reduced in most of the cities, it strengthened in Monterrey and Tijuana. In Monterrey, both in- and out-migration contributed to increasing the local female population. In Tijuana, only out-migration expanded the female population base, while in-migration had an increased masculinizing effect. For Monterrey, these patterns seem to reflect greater inflows of female migrants, partly reflecting women escaping from homicides of females in Ciudad Juárez. Monterrey also attracted women from cities on the Mexican–US border (including Tijuana) that were severely affected by the 2008 global financial crisis (Chavez et al. 2016). In Cuenca, internal migration had a strong masculinizing effect, but this decreased between 1995–2000 and 2005–10, reflecting a greater representation of males in out-migration flows. This may reflect economic restructuring towards female-dominated employment sectors, such as hospitality, accommodation, and trade, which reduced the labour demand for male workers and may have enticed women from neighbouring areas to migrate for employment (Chavez et al. 2016).

These findings point to a key historical change in patterns of internal migration in LA. In contrast to the dominant female selectivity in migration inflows to large LA cities during the 1930s to 1970s (Elizaga 1966; Herold 1979; Herrera 2013), a greater sex balance in these flows implies a reduction in this feminizing effect. The female selectivity during the 1930s to 1970s reflected the mass migration of women from rural areas to take low-skilled jobs in service activities in response to a shrinking agricultural sector (Useche 2013). As countries experienced rapid urbanization and agricultural decline, rural female migrants employed as domestic servants in high-income households were a common feature of the mobility system in LA countries up to the 1980s (Rodríguez and Busso 2009). Many rural towns have now evolved into small and intermediate urban areas and diversified their economies, developing tourism, accommodation, and food industry sectors, and expanding local job opportunities for women (Drentea 1998). Intermediate cities have also seen growing investment in university and vocational infrastructure, enlarging the range of local post-secondary education opportunities (Bulmer-Thomas 2003). Together, these developments may have promoted the retention of the local female population in small and intermediate cities, balancing the sex ratio in population movements to and from large LA cities.

Age structure

Table 5 and Figure 2 report the age CIMs. For the 1995–2000 period, they show negative CIMs for the populations aged 5–14 (children), 30–44 and 45–59 (working age), and 60+ (older people), indicating that internal migration reduced the share of these age groups in the local populations. There were large variations in the extent of these reductions. Internal migration appears to have generated the largest reductions in Panama City and Tijuana, as indicated by the relative CIM, leading to a reduction of nearly 5 per cent in the share of children in Panama City, and over 7 per cent in the share of older people in Tijuana. Reductions were marginal in Mexico City, with internal migration producing changes of less than 1 per cent. There were also exceptions to this downward trend. Internal migration acted to expand the 60+ population in Guadalajara and Guayaquil but these expansions were tiny.

The reductions in the shares of children and older people were reflected in a concomitant expansion in the share of the working-age population (i.e., ages 15–59). This was driven by a rising population aged 15–29 (see Figure 2), as internal migration acted to reduce the local populations at ages 30–44 and 45–59. Panama City and Tijuana experienced the largest increases in the share of working-age population, of 1.9 and 1.7 per cent, respectively, while Mexico City, Monterrey, Guadalajara, and Guayaquil saw the smallest increases (under 0.2 per cent). The biggest percentage increases can be observed in Tijuana, Quito, Cuenca, and Panama City, where internal migration operated to augment the local share of population aged 15–29 by over 4 per cent.

Table 5 reveals that in-migration was the main factor underpinning the reduction in the shares of people aged 5–14 and 60+ during 1995–2000, as indicated by consistently larger CIM^Is than corresponding CIM^Os. These negative CIM^Is indicate that the percentages of children and older people in in-migration flows to cities tended to be smaller than in the non-migrant population, and as a result, in-migration contributed to reduce the share of these groups in the local population. In Panama City, in-migration accounted for 93 per cent of the contraction in the share of the local population aged 5–14.

While out-migration tended to have a smaller impact than in-migration across all age groups and cities, we note that out-migration tended to have differentiated impacts: in some instances, amplifying the impacts of in-migration flows, and in others, counterbalancing these effects. The counterbalancing effects of out-migration were notable for older age groups (45–59 and 60+), as indicated by positive CIM^Os: out-migration tended to involve a smaller share of people aged 60+ compared with the non-migrant

Table 5 Compositional Impact of Migration (CIM): share of population aged 5–14, 15–29, 30–44, 45–59, and 60+, 1995–2000 and 2005–10

			Impact indicators: 1995–2000				Impact indicators: 2005–10			
	City	Country	CIM	Relative CIM (%)	CIM^I	CIM^O	CIM	Relative CIM (%)	CIM^I	CIM^O
Population aged 5–14	Panama City	Panama	−1.08	−4.99	−1.00	−0.07	−0.91	−4.47	−0.82	−0.09
	Mexico City	Mexico	−0.16	−0.75	−0.10	−0.07	−0.10	−0.55	−0.03	−0.07
	Monterrey	Mexico	−0.21	−0.97	−0.22	0.01	−0.12	−0.62	−0.13	0.01
	Guadalajara	Mexico	−0.18	−0.75	−0.14	−0.04	−0.25	−1.17	−0.22	−0.03
	Tijuana	Mexico	−0.73	−2.95	−0.62	−0.11	−0.15	−0.67	−0.07	−0.07
	Quito	Ecuador	−0.55	−2.49	−0.58	0.03	−0.34	−1.67	−0.29	−0.05
	Guayaquil	Ecuador	−0.19	−0.88	−0.31	0.11	−0.07	−0.34	−0.15	0.08
	Cuenca	Ecuador	−0.51	−2.14	−0.56	0.06	−0.38	−1.81	−0.41	0.03
Population aged 15–29	Panama City	Panama	1.49	4.91	1.60	−0.11	1.21	4.48	1.29	−0.08
	Mexico City	Mexico	0.51	1.57	0.54	−0.04	0.37	1.32	0.36	0.01
	Monterrey	Mexico	0.73	2.22	0.88	−0.15	0.57	2.06	0.78	−0.20
	Guadalajara	Mexico	0.39	1.18	0.47	−0.08	0.49	1.65	0.59	−0.10
	Tijuana	Mexico	2.32	7.08	2.43	−0.12	0.75	2.57	0.95	−0.20
	Quito	Ecuador	1.90	6.15	2.21	−0.31	1.32	4.41	1.68	−0.36
	Guayaquil	Ecuador	0.69	2.22	0.98	−0.29	0.40	1.37	0.65	−0.24
	Cuenca	Ecuador	1.83	5.72	2.14	−0.31	1.28	3.99	1.79	−0.51
Population aged 30–44	Panama City	Panama	0.22	0.89	0.23	−0.01	0.24	0.97	0.26	−0.02
	Mexico City	Mexico	−0.24	−0.94	−0.16	−0.08	−0.18	−0.70	−0.01	−0.17
	Monterrey	Mexico	−0.29	−1.17	−0.22	−0.07	−0.17	−0.66	−0.06	−0.12
	Guadalajara	Mexico	−0.20	−0.88	−0.03	−0.17	−0.23	−0.93	−0.02	−0.20
	Tijuana	Mexico	−0.67	−2.70	−0.61	−0.06	−0.35	−1.28	−0.18	−0.16
	Quito	Ecuador	−0.78	−3.23	−0.59	−0.19	−0.60	−2.52	−0.32	−0.28
	Guayaquil	Ecuador	−0.33	−1.38	−0.25	−0.08	−0.22	−0.94	−0.11	−0.11
	Cuenca	Ecuador	−0.56	−2.58	−0.44	−0.11	−0.46	−2.08	−0.25	−0.20
Population aged 45–59	Panama City	Panama	−0.39	−2.72	−0.50	0.12	−0.28	−1.71	−0.40	0.12
	Mexico City	Mexico	−0.08	−0.64	−0.18	0.10	−0.06	−0.33	−0.20	0.14
	Monterrey	Mexico	−0.14	−1.08	−0.26	0.12	−0.16	−0.99	−0.34	0.18
	Guadalajara	Mexico	−0.03	−0.24	−0.19	0.16	−0.03	−0.20	−0.20	0.17
	Tijuana	Mexico	−0.46	−4.14	−0.65	0.19	−0.20	−1.35	−0.45	0.25
	Quito	Ecuador	−0.38	−2.81	−0.63	0.25	−0.27	−1.74	−0.63	0.35
	Guayaquil	Ecuador	−0.17	−1.29	−0.29	0.12	−0.11	−0.67	−0.26	0.16
	Cuenca	Ecuador	−0.45	−3.57	−0.62	0.17	−0.28	−1.96	−0.62	0.34
Population aged 60+	Panama City	Panama	−0.25	−2.78	−0.32	0.07	−0.26	−2.35	−0.33	0.07
	Mexico City	Mexico	−0.02	−0.28	−0.10	0.08	−0.04	−0.34	−0.12	0.09
	Monterrey	Mexico	−0.09	−1.18	−0.18	0.09	−0.12	−1.18	−0.25	0.13
	Guadalajara	Mexico	0.03	0.34	−0.11	0.13	0.02	0.18	−0.15	0.17
	Tijuana	Mexico	−0.45	−7.07	−0.55	0.10	−0.06	−0.85	−0.25	0.19
	Quito	Ecuador	−0.19	−2.07	−0.41	0.22	−0.10	−1.01	−0.43	0.33
	Guayaquil	Ecuador	0.01	0.08	−0.13	0.14	0.00	−0.02	−0.12	0.12
	Cuenca	Ecuador	−0.32	−3.19	−0.52	0.20	−0.16	−1.51	−0.50	0.35

Note: The CIM statistic indicates the absolute percentage point change in the population shares measured at the census date and five years earlier; relative CIM shows the percentage change over the five-year period; and CIM^I and CIM^O quantify the impacts of in- and out-migration, respectively, on the population shares.
Source: Authors' calculations based on census data. Data covering the 1995–2000 and 2005–10 census periods were used for Panama and Mexico, and data covering the 1996–2001 and 2005–10 census periods were used for Ecuador.

population, amplifying the presence of this age group. This finding points to the fact that the overall impact of internal migration on reducing the local proportion of older people was not because of older people leaving large cities, but because of working-age people moving in, particularly those in the 15–29 age group. This result is consistent with observed low propensities to migrate among older people (Rogers et al. 1978), and also reflects the absence of a double bulge in the age distribution of migration in LA countries (ECLAC 2012). In industrialized countries, migration age profiles tend

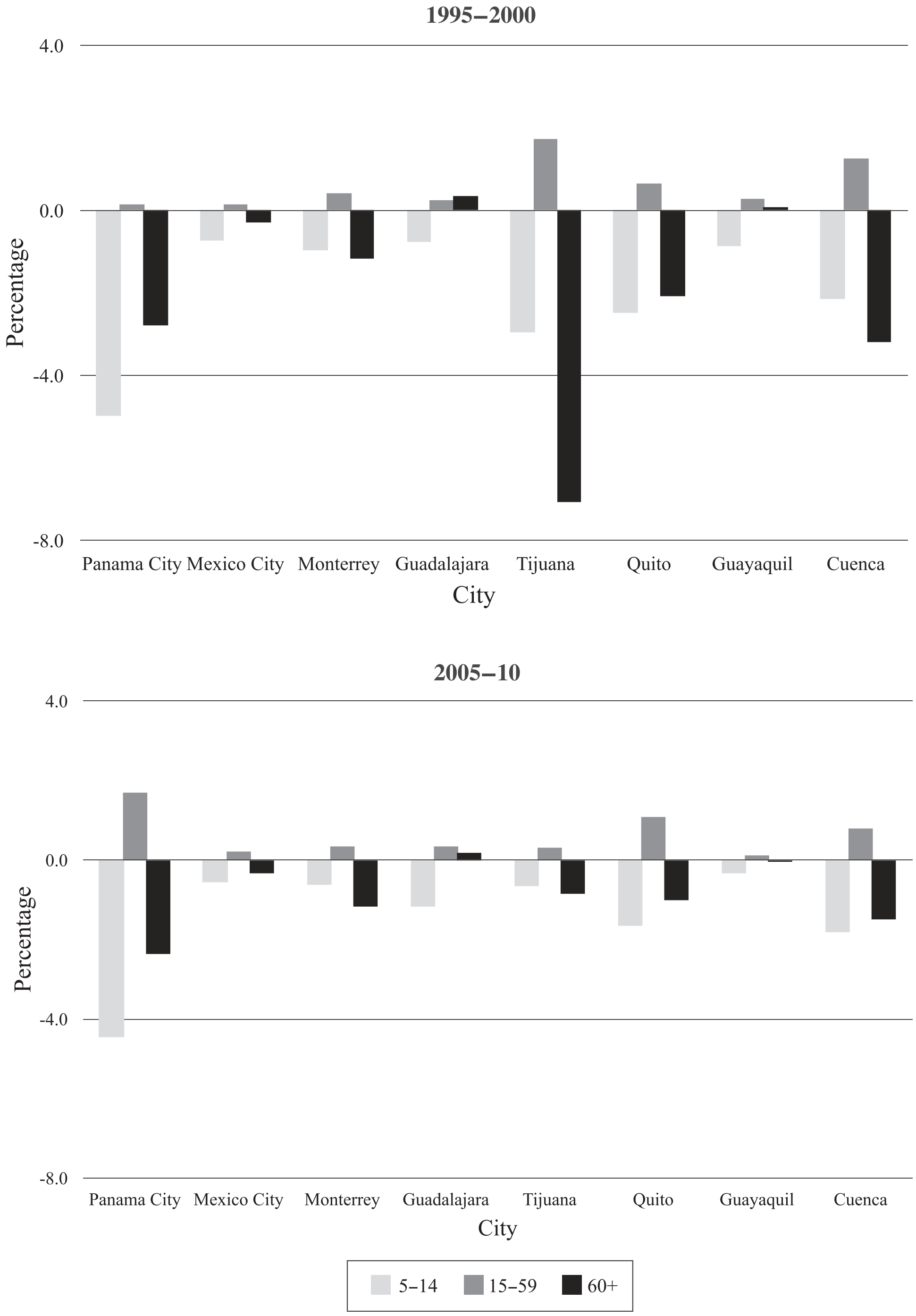

Figure 2 Relative Compositional Impact of Migration (CIM) index: share of population aged 5–14, 15–59, and 60+, eight Latin American cities, 1995–2000 and 2005–10

Source: Authors' calculations based on census data. Data covering the 1995–2000 and 2005–10 census periods were used for Panama and Mexico, and data covering the 1996–2001 and 2005–10 census periods were used for Ecuador.

to peak first around the mid-20s to early 30s, and peak for a second time around retirement age, reflecting the mobility of older people to rural and coastal areas (Hugo and Bell 1998). In LA countries, this second peak does not occur, as retirees living in large cities tend to stay put to take advantage of the high quality of local healthcare facilities (López-Calleja and Morejón 2015).

As in 1995–2000, Table 5 indicates that internal migration continued acting to reduce the shares of children and older people in Panama City and large Mexican and Ecuadorian cities during 2005–10. However, smaller CIMs in 2005–10 than in 1995–2000 reveal that this effect diminished in size. The average CIMs (not shown) for the shares of people aged 5–14 and 60+ across cities reduced by around

30 and 10 per cent, respectively, between 1995–2000 and 2005–10. Among the child population, this reduction was primarily due to a lessening impact of in-migration, while among older people it was driven by a strengthening of the impact of out-migration. For the former, this points to the fact that, as in 1995–2000, in-migrants continued to include a smaller share of people aged 5–14 than in the local non-migrant population, although this difference was less pronounced in 2005–10, generating a smaller reduction in the local child population. For the latter, this result indicates that the percentage of people aged 60+ in outflows from large cities decreased between 1995–2000 and 2005–10, with out-migration thus acting to increase the local share of older people.

Taken together, these patterns concurrently reflect an increasing share of people of working age, specifically young adults aged 15–29, moving into large Panamanian, Mexican, and Ecuadorian cities. This evidence is consistent with the pattern of young people in industrialized countries moving into large cities in pursuit of better education, employment opportunities, and a more vibrant lifestyle (Williamson 1988; Fielding 1992; Rowe, Corcoran, et al. 2017). However, the rejuvenating effect of internal migration on the age structure of cities in our sample declined and this seems to be linked to an overall reduction in the share of young migrants to large cities in response to the dispersal of employment opportunities, which expanded to medium-sized cities (Gilbert 1996). This decline in the rejuvenating effect of internal migration is expected to continue as LA countries move to more advanced stages of the demographic and urban transition, at which time an acceleration of population ageing and a strengthening of medium-sized cities' economies are expected (Gilbert 1996; Rodríguez 2011a).

Educational levels

In contrast to the impacts of internal migration on boosting the working-age population, examining the education CIMs for the population aged 25+ reveals that internal migration tended to have a downgrading effect on education by reducing the average years of schooling in local populations (Table 6 and Figure 3). The CIMs of 0.1 or less, however, reveal that this effect was marginal, indicating decreases of less than 0.1 years of schooling on average. While there was variation in the main source contributing to this effect across cities, out-migration appears to have consistently eroded the local human capital base, as indicated by negative CIM^{O}s, except for Panama City. In Ecuadorian cities, both in- and out-migration contributed to generate the largest reductions in the local average years of schooling. In Mexican cities, the *downgrading educational* effect produced by out-migration was offset by in-migration of people with higher education than the non-migrant population in both census periods.

These results are illuminating in two ways. First, they reveal a tendency for more educated people to leave major cities in Ecuador, Mexico, and Panama. This is contrary to what might be anticipated, as large cities concentrate educational and career development opportunities. This pattern appears to resemble the 'escalator region' process that characterizes the internal migration system in the UK (Fielding 1992), with young people moving to the South-East region of England to develop their careers, and then 'stepping off' the escalator and moving down the urban hierarchy once they achieve a high occupational status and family-related reasons gain importance. Second, the results reveal that in-migration into large cities does not necessarily lead to a reduction in the average years of schooling. We find that in-migration increased local levels of education in Mexican cities, in contrast to the traditional assumption that large cities attract a disproportionate number of less educated people to take up low-skilled jobs and education.

As pointed out earlier, a key assumption of our method is the time invariability of population attributes. As formal education tends to increase over time, we checked the robustness of our results by comparing the CIMs for the population aged 25+ with those for two older age groups: 30–44 and 45–49. The results were consistent. Although there were variations in magnitude, the CIMs for the 30–44 and 45–49 age groups were predominantly negative. They indicate that internal migration contributed to reduce the average years of schooling and that the out-migration of more educated people tended to be the main factor underpinning this reduction.

Additionally, Table 6 reveals that the downgrading effect of migration on education has weakened over time. In the 2005–10 period, while migration continued to reduce the average number of years of schooling, the associated reductions were 0.06 years or less. In part, this weakening in the overall impact of migration tended to reflect a lessening of the effects of both in- and out-migration on lowering the average years of schooling, but it may also relate to a decline in the propensity to move among highly educated people. This decline in migration probabilities is a common feature of mobility patterns in

Table 6 Compositional Impact of Migration (CIM): average years of schooling for population aged 25+, 30–44, and 45–49, 1995–2000 and 2005–10

			Impact indicators: 1995–2000				Impact indicators: 2005–10			
	Cities	Country	CIM	Relative CIM (%)	CIM^I	CIM^O	CIM	Relative CIM (%)	CIM^I	CIM^O
Population aged 25+	Panama City	Panama	−0.08	−0.79	−0.08	0.00	−0.04	−0.40	−0.05	0.01
	Mexico City	Mexico	−0.02	−0.28	0.00	−0.03	−0.02	−0.17	0.02	−0.03
	Monterrey	Mexico	−0.02	−0.25	0.03	−0.05	−0.02	−0.16	0.03	−0.04
	Guadalajara	Mexico	−0.01	−0.13	0.05	−0.06	−0.02	−0.20	0.05	−0.07
	Tijuana	Mexico	−0.03	−0.32	0.00	−0.03	0.00	0.02	0.01	−0.01
	Quito	Ecuador	−0.10	−0.96	−0.06	−0.03	−0.04	−0.39	−0.02	−0.03
	Guayaquil	Ecuador	−0.10	−1.13	−0.08	−0.02	−0.06	−0.67	−0.02	−0.05
	Cuenca	Ecuador	−0.08	−0.90	−0.05	−0.03	−0.06	−0.57	0.00	−0.06
Population aged 30–44	Panama City	Panama	−0.09	−0.82	−0.09	0.00	−0.07	−0.57	−0.08	0.01
	Mexico City	Mexico	−0.03	−0.29	0.00	−0.03	−0.02	−0.22	0.01	−0.03
	Monterrey	Mexico	−0.02	−0.24	0.02	−0.04	−0.01	−0.11	0.02	−0.03
	Guadalajara	Mexico	−0.01	−0.08	0.05	−0.05	−0.02	−0.18	0.05	−0.07
	Tijuana	Mexico	−0.05	−0.60	−0.03	−0.02	0.00	0.02	0.01	−0.01
	Quito	Ecuador	−0.10	−1.01	−0.06	−0.04	−0.03	−0.28	−0.03	0.00
	Guayaquil	Ecuador	−0.09	−0.96	−0.08	−0.02	−0.07	−0.67	−0.02	−0.05
	Cuenca	Ecuador	−0.11	−1.07	−0.10	−0.01	−0.05	−0.41	−0.03	−0.02
Population aged 45–49	Panama City	Panama	−0.08	−0.86	−0.09	0.01	−0.05	−0.44	−0.06	0.01
	Mexico City	Mexico	−0.01	−0.13	0.00	−0.01	−0.01	−0.11	0.01	−0.02
	Monterrey	Mexico	−0.01	−0.14	0.01	−0.02	−0.02	−0.19	0.00	−0.02
	Guadalajara	Mexico	0.00	−0.06	0.03	−0.03	0.00	0.05	0.02	−0.02
	Tijuana	Mexico	−0.11	−1.63	−0.10	−0.01	−0.02	−0.24	−0.03	0.01
	Quito	Ecuador	−0.10	−1.17	−0.08	−0.02	−0.03	−0.32	−0.03	−0.01
	Guayaquil	Ecuador	−0.08	−1.00	−0.08	0.00	−0.04	−0.47	−0.02	−0.03
	Cuenca	Ecuador	−0.06	−0.69	−0.07	0.01	−0.02	−0.18	−0.02	0.00

Note: The CIM statistic indicates the difference between the average number of schooling years measured at the census date and five years earlier; relative CIM shows the percentage change over the five-year period; and CIM^I and CIM^O quantify the impacts of in- and out-migration, respectively, on average years of schooling.
Source: Authors' calculations based on census data. Data covering the 1995–2000 and 2005–10 census periods were used for Panama and Mexico, and data covering the 1996–2001 and 2005–10 census periods were used for Ecuador.

many countries across the globe (Bell and Charles-Edwards 2013; Rowe 2018), and LA countries are no exception (Rodríguez 2011a). In Chile, Rowe (2013b) found that the odds ratio of migration for tertiary-educated individuals declined from 2.63 in the 1977–82 census period to 1.93 in the 1997–2002 period and that these individuals were more likely to live in the main metropolitan area of the country, Santiago. An increasing 'rootedness' of highly educated individuals, coupled to long-distance commuting, has been linked to this pattern (Cooke 2011; Rowe and Bell 2018) and it may explain the declining effects of migration on the educational composition of large LA cities.

National pictures

Thus far, we have discussed the results according to individual characteristics but what do they mean for the countries studied? This subsection links the observed patterns of internal migration impact to socio-economic and demographic processes in each national setting, acknowledging that a comprehensive analysis of the factors underpinning these processes is beyond the scope of this paper.

In Panama City, in-migration consistently operated to enlarge the population in the 15–29 and 30–44 age groups, reduced local educational levels, and transitioned from having a feminizing to a masculinizing effect on population composition. Over the last 25 years, Panama City has experienced rapid economic growth due to the expansion of trade, real estate, and professional service activities associated with the Panama Canal. These activities have attracted domestic and international migrants, and unlike in many LA nations, they form part of the formal economy (Chavez et al. 2016). They tend to employ a larger share of individuals aged 15–29 and 30–44 than aged 45–59, thus attracting a

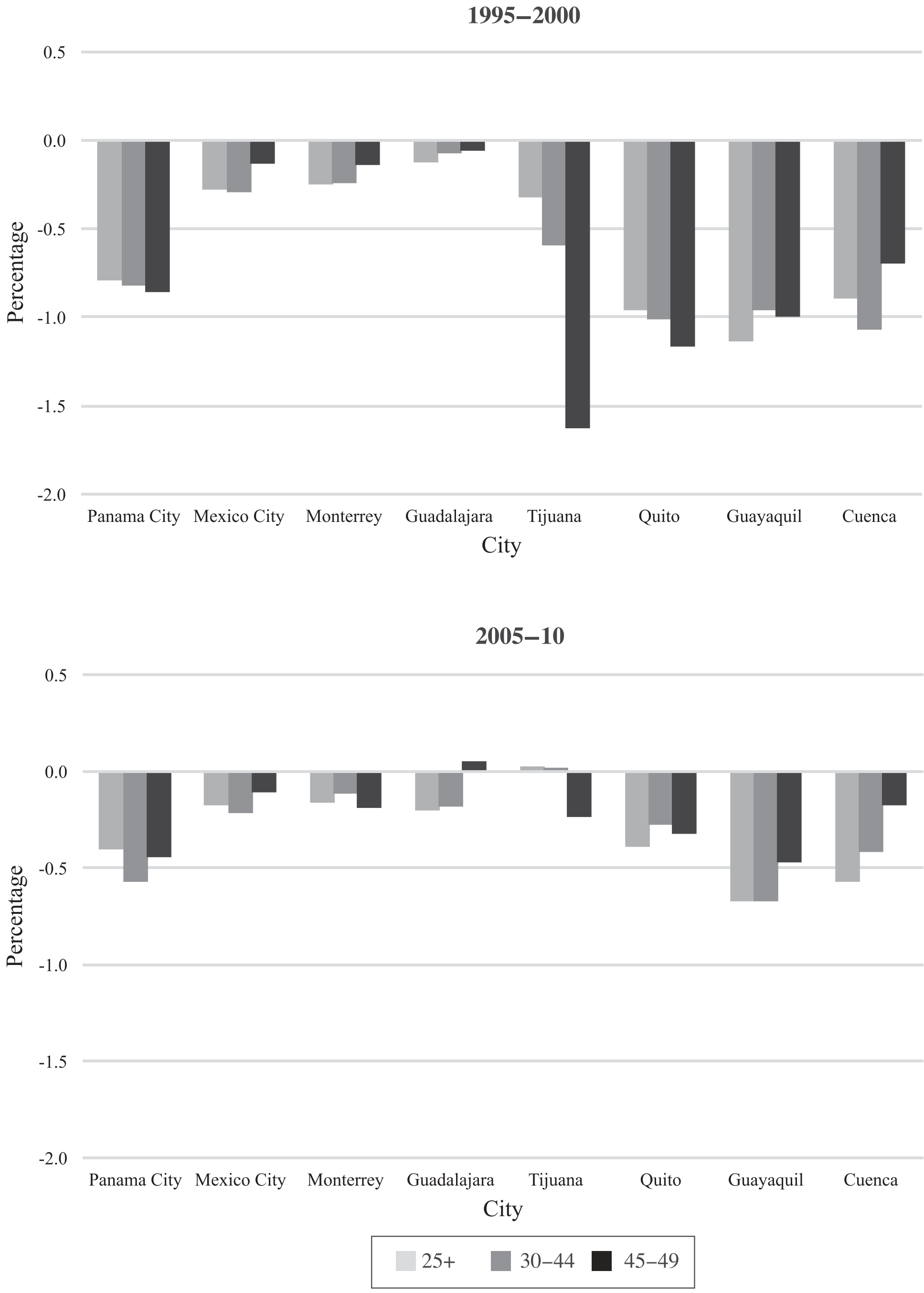

Figure 3 Relative Compositional Impact of Migration (CIM) index: average years of schooling for population aged 25+, 30–44, and 45–49, eight Latin American cities, 1995–2000 and 2005–10

Source: Authors' calculations based on census data. Data covering the 1995–2000 and 2005–10 census periods were used for Panama and Mexico, and data covering the 1996–2001 and 2005–10 census periods were used for Ecuador.

disproportionate percentage of non-local labour in the two former age groups (Chavez et al. 2016). These activities are also male dominated and this may explain the masculinization of migration flows to Panama City.

In Ecuador, a series of factors have contributed to shaping migration patterns during the end of the twentieth and start of the twenty-first centuries; these include a socio-economic crisis, dollarization of the local consumer market, and a larger public administration sector (Calderón et al. 2016). These factors encouraged international emigration, consolidated Quito (the capital) as the main national migration destination, and resulted in a positive net migration balance in Guayaquil. Concurrently, our results indicate that the impact of internal migration on the local population compositions by age and education was similar across large Ecuadorian cities:

increasing the local share of people aged 15–29 and reducing average years of schooling. However, its impact on the sex composition exhibited spatial variations, feminizing the local population in Quito and Guayaquil, but masculinizing that of Cuenca. The in-migration of male workers to replace local men migrating overseas appears to be underpinning this pattern in Cuenca (Herrera 2008).

In Mexico, as in other LA countries, the transition to an export-oriented economic development strategy has transformed patterns of internal migration, resulting in net migration losses in Mexico City (Useche 2013). Historically, Mexico City has been the most attractive destination for migrants but the development of new poles of economic activity, driven by the maquiladora sector, has promoted population movement to cities on the Mexican–US border (Gilbert 1996). In the maquiladora industry, young women are a desirable source of labour as they typically work longer hours than men and display a greater dexterity in performing repetitive assembly tasks (Chavez et al. 2016). Women are thus attracted to Mexican–US border cities for employment and this can be linked to an augmented *feminizing* effect of internal migration in Tijuana during the 2005–10 period—an effect that may be representative of Mexican–US border cities in general. This effect was also seen in Mexico City, Monterrey, and Guadalajara, where it was coupled with a consistent erosion of local educational levels (reduction in the average years of schooling), driven by the out-migration of people with higher educational qualifications than the locals (Table 6).

Conclusions

Despite its wide-ranging implications, little progress has been made in understanding and measuring the impacts of internal migration on the socio-demographic composition of local areas. A key obstacle has been the lack of an approach to capture the simultaneous influences of key population components that determine the spatial impact of internal migration: the size, balance, and selectivity of migrant and non-migrant populations. In this paper, we proposed a method that captures the dynamics of these components, exploits the availability of census data, and returns a single summary measure, the *Compositional Impact of Migration* (CIM). We showed how this index can be decomposed to measure the relative contributions of in- and out-migration.

We applied our method to quantify the impact of internal migration on the sex, educational, and age compositions of eight large LA cities. The estimated impacts were generally small, echoing the inertia of human settlement patterns in large cities, but systematic patterns emerged. Internal migration tended to reduce the local sex ratio and average years of schooling, and to increase the percentage of young adults (aged 15–29) in the local population, namely, having *feminizing*, *downgrading educational*, and *demographic window* effects on local population structures.

We also showed that the strength of these effects diminished between 1995–2000 and 2005–10. A greater representation of males and a smaller share of people aged 15–29 in the in-migration flows to these cities reduced the *feminizing* and *demographic window* effects, while a larger proportion of highly educated people moving in diminished the *downgrading educational* effect. We also highlighted that the estimated impacts of migration were small in terms of the aggregate impacts on metropolitan populations. They are likely to be more acute in particular zones within metropolitan regions, in places where net migration gains and losses are concentrated. Future research is required to determine the extent of these impacts at a sub-metropolitan scale.

It is important to recognize that our method focuses on quantifying the migration effect of time-invariant attributes, not time-varying characteristics. Our work contributes to advancing migration research in three ways. First, it provides tools to expand our knowledge of the impacts of internal migration. To date, existing scholarship has assessed the impacts of migration by examining flows and net migration balances. The proposed method enables us to complement this knowledge by quantifying and examining the compositional effects of migration on local populations.

Second, our work has the potential to guide policy design. Measurement and an understanding of the compositional impacts of internal migration are key inputs for policy development. Existing migration policies tend to focus on restricting in-migration (United Nations 2010). Yet, our findings revealed that internal migration may lead to a *demographic window* effect, expanding the local working-age population, which may lead local governments to promote policies that focus on in-migration. Concurrently, our findings revealed that migration tended to reduce local levels of education, which could justify restrictive in-migration policy measures. Yet, they also indicated that out-migration is the main mechanism for such loss in educational levels and, taken together, these results invite us to move away from traditional policy approaches focused on in-migration to adopt a more comprehensive framework that considers both in- and out-migration, as well as the size, balance, and composition of migration flows.

Third, our findings also have implications for further research on migration within LA countries. Prior work has documented a gradual reduction in female selectivity in the internal migration systems of LA countries (Rodríguez 2004), but to date, little is known about how this change has shaped metropolitan populations. Our results revealed that migration continues to have a feminizing effect on these population structures, but it is diminishing, and in certain cities—including Panama City, Guadalajara, and Cuenca—a masculinizing effect has emerged. This finding motivates a fruitful avenue of future research: to develop a better understanding of the socio-economic changes and explain the shifting sex selectivity of migration flows in LA countries. This would involve examining changes in the labour market of cities, specifically changes in the domestic sector, which has been a major employer of women historically, but has experienced a gradual downsizing (Chant 1999). Understanding the changing sex selectivity of migration would also require an investigation of the socio-economic changes in places of origin. In the 1970s and 1980s, these places were predominantly rural and highly dependent on agricultural activity. They are now small, vibrant urban environments, but little is known about their industrial structure or employment patterns.

Notes and acknowledgements

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2 We acknowledge the comments from the editor of *Population Studies* and two anonymous referees that helped to improve the quality of our manuscript.

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Appendix

Table A1 Factual Values (FV) and Counterfactual Values (CFV) for sex, age, and education, 1995–2000 and 2005–10

			Percentage by age group										Average years of schooling by age band					
	Sex ratio		Population aged 0–14		Population aged 15–29		Population aged 30–44		Population aged 45–59		Population aged 60+		Population aged 25+		Population aged 30–44		Population aged 45–59	
City	FV	CFV	FV	CFV	FV	CFV	FV	CFV	FV	CFV	FV	CFV	FV	CFV	FV	CFV	FV	CFV
1995–2000																		
Panama City	96.3	96.6	20.5	21.6	31.9	30.4	25.0	24.8	13.8	14.2	8.7	9.0	10.1	10.2	10.9	11.0	9.5	9.6
Mexico City	91.9	92.4	21.9	22.1	32.6	32.1	24.7	25.0	12.9	13.0	7.8	7.8	8.8	8.8	9.6	9.6	7.8	7.8
Monterrey	97.5	97.8	21.5	21.7	33.7	32.9	24.5	24.8	12.7	12.8	7.7	7.8	8.9	8.9	10.0	10.0	7.8	7.8
Guadalajara	92.8	92.9	24.5	24.6	33.3	32.9	22.6	22.8	12.1	12.1	7.6	7.5	8.3	8.3	9.2	9.2	7.3	7.3
Tijuana	100.4	100.7	24.1	24.8	35.0	32.7	24.3	25.0	10.6	11.1	5.9	6.4	7.9	7.9	8.6	8.7	6.6	6.7
Quito	92.6	93.3	21.5	22.1	32.8	30.9	23.5	24.2	13.0	13.4	9.2	9.4	9.8	9.9	9.5	9.6	8.1	8.2
Guayaquil	95.3	96.0	22.0	22.2	31.7	31.0	23.8	24.2	12.9	13.1	9.5	9.5	8.7	8.8	9.7	9.8	8.1	8.2
Cuenca	87.7	86.7	23.2	23.7	33.9	32.1	21.1	21.6	12.1	12.6	9.7	10.0	8.9	9.0	10.2	10.3	8.2	8.2
2005–10																		
Panama City	96.4	96.4	19.5	20.4	28.3	27.1	26.0	25.0	16.3	16.6	10.7	11.0	11.1	11.1	11.8	11.8	11.1	11.1
Mexico City	91.7	92.0	18.4	18.5	28.5	28.1	25.1	25.3	17.3	17.3	10.8	10.8	9.7	9.8	10.7	10.7	9.4	9.4
Monterrey	98.1	98.7	19.8	20.0	28.5	27.9	25.8	26.0	16.1	16.3	9.8	9.9	9.9	9.9	10.7	10.7	9.9	9.9
Guadalajara	92.0	91.8	21.4	21.5	30.2	29.7	24.0	24.3	14.8	14.8	9.6	9.6	9.4	9.5	10.3	10.3	9.2	9.2
Tijuana	97.7	98.3	22.0	22.2	30.0	29.2	26.9	27.3	14.2	14.4	6.9	6.9	9.1	9.1	9.8	9.8	8.5	8.5
Quito	93.2	93.9	20.1	20.4	31.2	29.9	23.2	23.8	15.5	15.7	10.1	10.2	10.7	10.8	11.5	11.5	10.6	10.7
Guayaquil	96.5	96.8	21.6	21.7	29.9	29.5	23.5	23.7	15.8	15.9	9.2	9.2	9.5	9.6	10.2	10.3	9.4	9.4
Cuenca	89.0	88.7	20.7	21.1	33.3	32.1	21.5	21.9	14.2	14.5	10.3	10.4	10.1	10.1	11.1	11.2	10.0	10.0

Note: FV and CFV correspond to the Factual and Counterfactual Values, respectively, as described in the 'Method' subsection. The FV indicates the score of a statistic at the census date and the CFV indicates the score of the same statistic measured five years ago.
Source: Authors' calculations based on census data.